Excise taxes, tax incidence and the flight to quality: Evidence from scanner data

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Abstract

Because excise taxes are independent of product price, rate hikes are predicted to lower the relative cost of high-priced goods, encouraging a shift towards their purchase. Using scanner data on cigarette sales from 29 states over a six-year period, we examine whether tax hikes encourage a flight to quality. Results demonstrate that a 1¢ hike in taxes increases retail prices of name brand and generic cigarettes by exactly 1¢. We find no tax-induced substitution towards name brands but a large substitution away from carton to pack sales, suggesting that tax hikes encourage within-product changes in purchase but little between-product substitution.

JEL classification: H22

Introduction

Excise taxes are typically levied on the purchase or sale of a product with the tax value being independent of product price. For this reason, excise taxes are often referred to as *quantity* taxes. In markets with homogenous products a rise in the excise tax will impact all products the same but the outcome is more complicated when taxed products vary along a number of dimensions. An excise tax on a commodity that has multiple characteristics may result in a substitution effect where consumers shift away from the product's characteristics that are taxed towards other characteristics that go untaxed. The basic theory outlining the impact of excise taxes for multi-attribute goods was first proposed by Barzel (1976) who extended the basic models of Lancaster (1966). In the Barzel model, a tax on quantity leads to a lower opportunity cost of purchasing quality which results in a "flight to quality" by consumers.[1] Quality, more generally, represents all other characteristics of a good that are not measured by quantity.[2]

Barzel tested the model using price data from the cigarette industry exploiting state-level changes in the tax rate. Barzel finds that a 1¢ increase in taxes leads to a greater than 1¢ increase in state-level average retail prices, which he interprets as being driven by a tax-induced increase in the market share of high-priced cigarettes.

In this paper we return to the original example in Barzel (1976) and consider whether higher tax rates induce a shift to higher quality cigarettes. The cigarette industry is segmented along price lines with most cigarettes falling into either a name-brand or a low-cost (generic) category. Unlike previous papers on this subject, we utilize high-frequency price and quantity

[1] The basic result in Barzel is a formalization of a conjecture by Alchian and Allen (1964) that a per-unit transaction cost such as transportation reduces the relative price of higher quality goods and encourages the trade of higher quality products.
[2] Following Barzel (1976) are other theoretical works that address the issue of taxation and changes to product quality. Examples include Leffler (1982), Kay (1991), Bohanon and Van Cott (1991), Goolsbee (2004), Goodhue, and LaFrance, and Simon (2009).

data available from supermarket scanners. Our sample includes data on weekly sales of all cigarettes sold in 812 stores from 43 markets covering 29 states over the 2001-2006 time period. The time period is excellent for our purpose because there were 32 state excise tax hikes and one tax decline, with many of the tax hikes being large changes, making it easy to detect the price and quality impacts of this policy lever.

We aggregate the data into cigarette type/state/month cells allowing us to initially examine the impact of taxes on generic and name-brand prices. Given the high-frequency data, we observe that retail prices adjust immediately to excise tax hikes. Our longitudinal model with both state and time fixed-effects show that for each 1¢ increase in excise taxes, retail prices of name-brand and generic cigarettes increase by exactly 1¢. Because generics are about 20 percent cheaper than name-brand cigarettes, tax hikes therefore increase the relative cost of generics compared to their higher-priced counterparts. If the Barzel model is correct, we should see tax hikes producing a flight to quality and a decline in the generic market share. In contrast to previous work, we find little evidence of this inter-brand substitution. Our estimates of the impact of taxes on generic market share are quantitatively small and statistically insignificant.[3]

We do however find a qualitatively large shift along a different dimension of quality. Cigarettes are sold in different packaging and in different places with per unit prices varying considerably based on the type of purchase. For example, most cigarette purchases are either by pack (20 cigarettes) or carton (10 packs). On a per pack price basis, carton purchases are about 12 percent cheaper than single packs. If the convenience of purchasing single packs (e.g., easier to store, more convenient to carry around) is an untaxed quality dimension, then we should see a shift to single pack purchases. As was the case for generics and name-brands, we find that per

[3] Throughout the paper, we will use the words "generic" and "discount" cigarettes interchangeably and in general, we use these phrases as the complements of "name-brand" cigarettes. As we note below, we use how various online smoke shops catalog cigarettes to define whether a cigarette is name-brand or not.

pack retail prices of both cartons and single pack purchases increase penny for penny given excise tax hikes, meaning that tax increases lower the relative cost of single pack purchases. In this case, we find a quantitatively large flight to quality in that for every 50¢ increase in cigarette taxes, which is in real December 2006 prices, close to the median tax change in our sample, carton market shares fall by 3.1 percentage points. Analysis of sales by brand suggests all of this effect is the movement from carton to pack sales within brand rather than across brands.

Our paper contributes to two distinct literatures. The first is a literature that examines the incidence of sales and excises taxes in general and taxes on cigarettes in particular. Authors have examined how taxes have impacted retail prices for a variety of products including gasoline (Chouinard and Perloff, 2004; Alm et al., 2009), alcohol (Kenkel, 2005; Young and Agnieszka-Kwapisz, 2002), imported television sets (Karp and Perloff, 1989) and a variety of products sold at retail establishments (Poterba, 1996; Besley and Rosen, 1999). This topic has however received the most attention in the case of cigarette excises taxes with a number of authors have examined how retail prices are impacted by changes in cigarette taxes. Much of this literature uses aggregate price data that varies at the state/year level. Studies using single prices within a state/year cell typically find that prices increase more than a penny for every cent increase in excises taxes (e.g., Sumner and Ward, 1981; Sullivan, 1985; Keeler et al., 1996; Evans and Farrelly, 1998; Evans and Ringel, 1999) although some find a one-for-one the tradeoff (Evans et al., 1999). Estimates based on a single average price per jurisdiction are potentially subject to the Barzel criticism that the tax effect is mixing tax incidence with quality changes. As a result, in recent years, a number of working papers have exploited data sets with heterogeneous prices and product quality to separate the quality-mix effect from tax incidence.

Hanson and Sullivan (2009) surveyed stores in Wisconsin before and after a tax hike and obtained price per store for a representative name-brand and generic cigarette. In this case, the authors found that taxes are over-shifted to consumers but the price effect is substantially lower for stores close to borders of states with lower-priced cigarettes. These authors did not examine actual transaction prices which may be important because coupon and in-store discounts are two frequent advertising devices in this industry.

Chiuo and Muehlegger (2010) used scanner data from a single chain of supermarkets in the Chicago area to examine the price impact of higher taxes on cigarettes from three manufacturers (Lorriard, Liggett and R.J. Reynolds). They estimated a price pass-through rate of just under one-for-one but like Hanson and Sullivan (2009) provided some evidence that stores closer to lower-taxed jurisdictions had lower pass-through rates.

DeCicca et al. (2010) used individual-level data on smokers from two rounds of the Current Population Survey's Tobacco Use Surveys where respondents provided the price paid for their last pack of cigarettes purchased, including any discounts or coupons. In their full sample results, the authors found that taxes increased penny for penny with excise tax hikes. The authors found little heterogeneity in the tax pass-through rate by smoking behavior such as smoking intensity, past quitting behavior, of cigarette type.

Harding et al. (2010) examined the tax incidence of higher cigarette taxes using individual-level data from the Nielsen Homescan data base. In this survey households scan the UPC code of all recent purchases. This paper analyzed cigarette purchases from roughly 160,000 transactions over the 2006/2007 period. In models with individual-level controls and UPC fixed effects, the authors found that 86 percent of a tax hike is passed on in the form of higher prices. The sample has the advantage of having individual level controls but survey respondents with purchases from certain stores do not record prices. Rather, the weekly price from the store is added to the data base, which may potentially overstate price when coupon sales are important.

Our study is in some respects broader than the work of Hanson and Sullivan (2009) and Chiuo and Muehlegger (2010) in that we include data from 29 jurisdictions and the data represent all cigarettes sold in supermarkets, not just a few brands. Our work also exploits to a greater degree than DeCicca et al. (2010) the timing of tax changes since they have data at roughly two different points in time. Our sample also has actual transaction prices rather than what consumers recall as the price. We also have price data from all purchases not just the last one made by a smoker. Our study is however more limited than Hanson and Sullivan (2009) and Harding et al. (2010) in that we only have sales data from supermarkets. Unlike Hanson and Sullivan, DeCicca et al., and Harding et al., we include actual transaction prices. Maybe most importantly, our analysis sample spans many tax hikes at a high frequency.

Our paper also contributes to the literature about whether excise taxes impact the quality of products purchased. Authors have examined whether excise taxes impact product quality for a number of products including gasoline (Nesbit, 2007), wine (James and Alston, 2002), capital investments (Goolsbee, 2004) and the purchase of football tickets (Bertonazzi et al., 1993). This topic has however received the most attention within the context of the generic cigarette market. Early work by Barzel (1976) and Johnson (1978) using aggregate prices find excise taxes increase prices by more than the tax hike, a result consistent with the original Barzel model. Sumner and Ward (1981) question this basic methodology suggesting that firms experience changes in factor prices more frequently than they are willing to respond with adjustments to product price. In the face of increasing factor prices due to inflation, for example, firms may choose to withhold product price hikes until some future point in time. An increase in an excise tax, then, is a coordinating event where backlogged product price adjustments are added to the price response that is directly related to the tax. Sumner and Ward find corroborating empirical evidence for their conjecture, and in doing so, they raise concern over the reliability of past empirical tests using one price as the outcome of interest.

In a more direct test of the flight to quality hypothesis, Sobel and Garrett (1997) measures changes in market shares of premium cigarettes resulting from excise taxes. Using annual generic market share data at the state level from 1990-94, Sobel and Garrett found that a 3¢ excise tax leads to a 1 percentage point increase in market share of premium cigarettes which is a large gradient. Our results are in stark contrast to these previous estimates in that we find no evidence of a tax-induced shift towards generics in this market.

Our results illustrate a number of interesting methodological, economic and policy points. The first is the benefit of high frequency and high quality scanner data to examine this question. Second, our results suggest that tax changes seem to induce little between-brand substitution but substantial within-brand differences in purchase behavior. So, although the predictions of the Barzel model were falsified in one dimension, they appear relevant in another. Third, from a policy perspective, our results indicate that in this market, retail prices are raised penny for penny with state tax hikes.

Data

The primary data set for this project is the IRI Academic Data Set that includes weekly store-level data on product sales, pricing, and promotion for 30 product categories in 48 U.S. Markets. The data covers the 2001-2006 time period and price and quantity data are collected at the Universal Product Code (UPC) level. The sample frame includes chain grocery stores and drug stores that supply data to Information Resources, Inc (IRI). The data set is described in detail in Bronnenberg et al. (2008) and applications for the data can be found at http://symphonyiri.com/Academics/tabid/172/Default.aspx. The data is available for academic researchers at a modest one-time cost of $500.

Markets are defined by geographic area and include counties associated with major metropolitan areas. Store names and locations are not provided but each store has a unique identification code that allows one to track within-store purchases over time. The data also provide information on when stores opened or started to supply data to IRI and when they closed or quit providing data to IRI. We restricted the sample to include only those stores that were in the sample continuously during the entire 2001 – 2006 period. Given the limited geographic information about stores, we also delete data from some areas because the markets may span multiple states. For this reason, we exclude data from the New York City, Chicago, New England, New Mexico/West Texas, and Washington, DC markets, leaving us with data from 43 markets from 29 different states. A list of states, markets within states and stores per market is displayed in Table 1.

Given the states in our sample, there are 33 separate state tax changes that allow us to identify the econometric model. A list of these tax changes, the effective date of the changes and the nominal value of the change in cents/pack is reported in Table 2. The data for this table is taken from the website of the Campaign for Tobacco Free Children.[4] The variation in the size of these tax changes ranges from a 10¢-drop in the per-pack tax in Oregon in 2004 to a high of 82¢-rise per pack in Arizona in 2006. The tax hikes are on average very large with the mean value being 42¢/pack, the median value being 39¢/pack, and 12 of the tax changes being 60¢ or larger

[4] http://tobaccofreekids.org/research/factsheets/pdf/0275.pdf.

per pack. Along with these state tax hikes, the federal tax on cigarettes was raised once during this period, rising from 34 to 39¢/pack on January 1, 2002.[5]

A limitation of the IRI data is that it only includes information from chain supermarkets and drug stores, retailers that are responsible for a minority of cigarette sales. In Table 3, we report data from the 2002 Census of Retail trade that provide information on total sales by detailed product line by retail type. The Census of Retail only provides data on aggregate tobacco sales but cigarettes represent roughly 92 percent of retail in tobacco.[6] In the table, we include all those retail sources with at least $1 billion in tobacco sales, which is 98 percent of all tobacco sales. Notice that the largest sources of sales are gas stations with convenience store which represents 41 percent of cigarette sales. Sales from supermarkets, drug stores and general merchandise stores (which include superstores selling groceries) account for 32.1 percent of cigarette sales. In our case, we must sacrifice market coverage to obtain the excellent detail in price and quantity data inherent in most scanner data sets. However, Hanson and Sullivan (2009) found that the price effect of higher taxes was the same for grocery stores, convenience stores, and tobacco stores or chains. This suggests that our reliance on a supermarkets and drugs stores should not reduce the generality of the study and its findings to the broader retail market for cigarettes.

Products are identified by a UPC code which is a unique 13-digit number that identifies the product and the manufacturer. The IRI data resources include a UPC lookup file that has detailed product characteristics including how many cigarettes are in a pack and whether the item is a single pack or part of a multipack purchase. We restricted our attention to cigarettes sold as part of 20 cigarette packs or cartons (i.e., 10 packs of 20 cigarettes/pack), eliminating

[5] http://www.cdc.gov/mmwr/preview/mmwrhtml/mm5819a2.htm.
[6] http://www.ers.usda.gov/Briefing/Tobacco/Data/table21.pdf.

promotional packages containing only two or three cigarettes. These restrictions eliminated less than 1 percent of cigarette sales in our sample.

For all cigarettes in our sample, we grouped cigarettes into either a name-brand or a discount (generic) group. Many online cigarette retailers group their products in this fashion and we initially used breakdowns from many of these internet sites to construct the groupings. We verified our groupings by looking at the average real pre-tax retail price and any obvious errors were corrected. In Table 4, we report our breakdown of some of the more well-known name-brand and discount cigarettes in our sample and the product's manufacturer. The name-brand group includes classic brands such as Marlboro, Winston, Salem and Camel, and the discount group includes brands like Basic, Doral, Tahoe, Malibu, and USA Gold.

The price and quantity data is aggregated to the year/store/week level with the first week starting January 1, 2001. We grouped the weeks into months with the month determined by the month of the last day of the week. Because some tax changes occur towards the middle of a month, monthly aggregates will mean that some months contain data from before and after tax hikes (see Table 2 for examples). We identify these "multi-tax" months in the data and expect that our estimates for these months and months when a single tax is paid will differ.

Given the mapping of stores into states, UPCs into product type (discount or name-brand) and weeks into month, we aggregate the data into type/state/year/month cells. Since there are 2 cigarette types (name-brand and discount), 29 states, 12 months per year and 6 years, there are a total of 4,176 observations in our primary analysis sample. To this data set we merge the nominal state and federal tax by month and the Consumer Price Index for All Urban Consumers.

Descriptive statistics for the analysis sample are included in Table 5. In the table, we provide the real retail price per pack, the real per-pack taxes (state and federal), and the real before-tax price (retail price less taxes). All means are weighted using the number of packs in the cell as the sample weight. All prices are in real December 2006 dollars, the last month in our sample. The table indicates that the name-brand cigarettes sold for $3.74/pack which is 61¢/pack higher than the average price for discount brands, an almost 20 percent price differential. During our period of analysis, about 28 percent of the retail price of name-brand cigarettes is federal and state excise taxes while this same number for discount brands is 37 percent. During our period of analysis, discount brands sold in supermarkets and drug stores represent 18.3 percent of the market. We will discuss the numbers in the bottom half of Table 5 in a moment.

Econometric model

To examine whether rising excise taxes generate a flight to quality, we exploit the panel nature of our data and use the within-market variation in taxes over time to identify our model. Initially, we consider the impact of higher excise taxes on the retail prices of name-brand and discount cigarettes. For this model, let P_{ist} be the average price for cigarettes of type i (either name-brand or discount) in state s in time period t. The basic analysis of covariance specification is given as

$$(1) \quad P_{ist} = \beta_0 + Tax_{st}\beta_1 + Multitax_{st} * Tax_{st}\beta_2 + Discount_i\beta_3 + \lambda_{1s} + \phi_{1t} + \varepsilon_{1ist}$$

where Tax_{st} is the real tax (state and federal) on cigarettes in state s for month t. Taxes and prices are measured in real December 2006 cents per pack. The coefficient of interest is β_1 which measures how much of the tax is passed onto consumers in the form of higher prices. We define the tax to be the one in effect on the last day of the month. As mentioned earlier, there may be some months that contain data from two different tax levels because of the timing of the tax adjustment. We control for this by including a separate tax coefficient for these transition months. The variable $Multitax_{st}$ is a dummy that equals 1 in months that contain data with two excise tax rates, and we expect $\beta_2 < 0$ because of the attenuation bias associated with these months. The variable $Discount_i$ is a dummy variable that equals 1 for discount brands and zero otherwise. The model incorporates a four-part error structure as we control for permanent differences in prices across cigarette type ($Discount_i$), states (λ_{1s}), months (ϕ_{1t}), plus any idiosyncratic differences in price that vary across all three of these dimensions (ε_{1ist}).

We can allow the impact of the taxes to vary by cigarette type which changes equation (1) to read

$$(2)\quad P_{ist} = \beta_0 + Name_i * Tax_{st}\beta_1^n + Discount_i * Tax_{st}\beta_1^d + Name_i * Multitax_{st} * Tax_{st}\beta_2^n + Discount_i * Multitax_{st} * Tax_{st}\beta_2^d + Discount_i\beta_3 + \lambda_{1s} + \phi_{1t} + \varepsilon_{1ist}$$

where $Name_i$ equals 1 for names brands and zero otherwise. This equationallows us to easily test the null hypothesis that the tax-induced change in retail prices is the same for both name, that is $H_o : \beta_1^n = \beta_1^d$.

Once we establish the impact of taxes on prices, we then investigate whether excise taxes encourage a flight to quality. In this case, we aggregate the data by market and the dependent variable is measured as MS_{st}^d which is the market share for discount brands in state s in year t. The estimating equation is similar to equation (1)

(3) $MS^d_{st} = \alpha_0 + Tax_{st}\alpha_1 + Multitax_{st} * Tax_{st}\alpha_2 + \lambda_{2s} + \phi_{2t} + \varepsilon_{2st}$

where all terms in the equation are defined above. As above, we control for fixed market and year effects and any impact of taxes on the relative demand for name and discount brands will be captured by the coefficient α_1. Given the high frequency of the data, we anticipate some autocorrelation within a market over time. As a result, in all models, we allow for arbitrary autocorrelation in error within a market over time.

Excise Taxes and Retail Prices for Generic and Name-Brand Cigarettes

Initially, we document the impact of excise taxes on retail prices for both name-brand and generic cigarettes. Before we report estimates for equations (1) and (2), we graphically illustrate the time series pattern of retail prices before and after tax hikes in six states that experienced large tax hikes during our period of analysis: Arizona, Massachusetts, Ohio, Washington, Pennsylvania and Connecticut. At the top of each graph, we note the date of the tax change and the nominal change in the state excise tax rate. For each graph, the vertical axis represents the change in nominal prices compared to the last full month before the tax change. The solid lines represent the average retail price for name-brands while the dotted lines provide data for generics.

Three results are clear from this graphical presentation. First, retail prices adjust immediately to the excise tax changes for both types of cigarettes with increases in the month taxes are raised. Second, the retail price change associated with tax hikes is on average about the same for both generics and name-brands. Note that in Ohio and Washington, the retail price change was the same for both groups, in Arizona and Pennsylvania generic prices rose slightly more than for name-brands while in Massachusetts and Connecticut the reverse was the case. Third, in all six cases, the post-tax hike increase in nominal prices was about equal to the nominal change in the excise tax. Although these graphics represent only 6 of the 33 tax hikes in our sample, most of the tax hikes in this period show similar patterns. Overall, these results are suggestive that in our regression estimates of equations (1) and (2), we will find that excise taxes increase retail prices about penny for penny and there will be little difference in the estimates for generic and name-brands.

In the first two columns of Table 6 we report estimates for equations (1) and (2) using these 4176 observations. The outcome variable is the real retail price (dollars per pack) deflated by the CPI to December 2006 prices. In these two regressions we add state and month/year effects and a dummy variable for whether the observation is for generic cigarettes. The key covariate is the real excise tax (state and federal) again in dollars per pack. We report parameter estimates, and in parentheses we report the standard errors that allow for arbitrary correlation in errors within a state over time. For all coefficients on the excise tax coefficients we also report, in square brackets, the p-value on the test for the null hypothesis that the coefficient equals 1.

Note that the model explains 98 percent of the variation in average prices across states and time and discount cigarettes are on average about 71 to 74 cents cheaper than name-brand cigarettes. In column (1) the coefficient on real taxes per pack is 1.03 meaning that average retail prices increase by almost exactly a dollar for a dollar increase in taxes. The p-value on the null hypothesis that this coefficient equals 1 is 0.70. In column (2) where we allow the coefficient on the excise tax variable to vary for name-brands and discounts, both coefficients equal 1.0 when rounded to the first decimal point. The p-values on the test that each parameter separately equals 1, both parameters are equal, or both parameters jointly equal 1 are very high and indicate we cannot reject any of these null hypotheses.

Overall, the results from columns (1) and (2) of Table 6 provide numeric equivalents to the visual evidence from Figure 1. Excise taxes raise retail prices for discount and name-brand cigarettes both by the amount of the tax. Given the large price difference between these two types of cigarettes, excise taxes will then increase the relative price of discount cigarettes compared to name-brand. To put some perspective on these numbers, suppose the average price of name-brand cigarettes is the value from Table 4 ($3.74) and state excise taxes are raised by 50 cents per pack, which is close to the median value in real terms of the tax hikes listed in Table 2. In this case, name-brand cigarette prices will increase by 13.4 percent, and assuming discounts are 73 cents cheaper per pack, their average price will increase by 16.6 percent.

Because the covariate of interest varies at the state level, there is a potential that the errors within a state are correlated over time. To guard against a high Type I error rate, we follow the recommendation of Bertrand et al. (2004) and calculate standard errors allowing for arbitrary correlation in within-state residuals. The asymptotic properties of this procedure have been demonstrated for when the number of groups approaches infinity and Monte Carlo evidence suggests that the estimates have a high Type I error rate when the number of groups is small (Wooldridge, 2005). We have 29 states so there may be a concern that the standard errors are arbitrarily low. In this case, the modest number of states does not appear to be a concern. Using a wild bootstrap procedure outlined by Cameron et al. (2008) that has been demonstrated to have good Monte Carlo properties even in cases with small numbers of groups, we estimate the p-values for some of the hypothesis tests from Table 6 and there is little of any difference in the p-values. For example, in column 1, the p-values on the test that the tax coefficient is zero in the clustered standard error case and via the wild bootstrap procedure are both 0 out to 6 decimal places. The p-value on the test that the tax coefficient in column (1) is 1 is 0.70 with clustered standard errors and 0.73 with the wild bootstrap procedure.

As expected, excise taxes increase the cost of discounts compared to name-brands. However, there is little evidence of a flight to quality in this context. In Figure 2, we present graphical evidence that suggests higher taxes generate a shift to discount brands. For the same six states considered in Figure 1, we graph the discount market share before and after large tax hikes. The solid line is the market share in a month while the vertical line represents the month the higher tax goes into effect. The dotted lines represent the pre- and post-tax hike average market shares. In four of the six states, there was an increase in discount market share (Arizona, Ohio, Washington and Connecticut) while there was a flight to quality in two states (Massachusetts and Pennsylvania). In most instances the tax-induced change in market share is small except for Pennsylvania which had a large discount market and experienced about a 10 percentage point drop in the discount market share.

The visual evidence is confirmed in the regression analysis reported in the final column of Table 6 where we present estimates of the market share model outlined in equation (3). A regression of the discount brand market share on market and time effects plus the real tax per pack produces a positive tax coefficient of 0.02 and a t-statistic of 1.86. The estimate suggests that at a p-value of 0.10 we can reject the null that taxes have no effect on discount market shares, and the effect is positive, albeit small. A 50¢ increase in the per-pack tax, which is a large change in price, is estimated to increase discount market shares by just over 1 percentage point. Overall, the evidence in Table 6 suggests there is no flight to quality and if anything, it appears that the generic/discount market increased slightly given these large tax changes.

One key result from Table 6 is that our estimates suggest that the entire tax hike is passed onto consumers in the form of higher prices. The model in Table 6 is however limited in that we aggregate data to the group level (generic or name-brand) within a state/month cell. Our model may generate biased estimates if tax incidence is correlated with some other cigarette characteristics (e.g., menthol or not, low tar/nicotine, manufacturer, etc.) that is not captured in the aggregate model but is correlated with demand elasticities and hence tax incidence. To examine this in more detail, we consider a different model where we aggregate data to the UPC/state/month level. In this case we estimate a model of the form

$$(4) \quad P_{ust} = \beta_0 + Tax_{st}\beta_1 + Multitax_{st} * Tax_{st}\beta_2 + \theta_{1u} + \lambda_{1s} + \phi_{1t} + \varepsilon_{1ust}$$

where u represents the unique UPC code. In this model, we include fixed time and state effects as well as a complete set of effects for all unique UPC codes. The variable ε_{1ust} represents an idiosyncratic error.

This specification expands the sample considerably. We have a total of 2,843 unique UPC codes and over 1.1 million UPC/state/month observations. Not all UPCs are sold in all states in all months so this is an unbalanced panel data set.

Results for equation (4) are reported in Table 7. In the first column, we report estimates using all 1.1 million observations, including state, month and year effects but excluding the UPC effects. The R^2 of this model is only 0.69 but we obtain an estimate on the tax coefficient of 0.99 and cannot reject the null that the coefficient is 1 given the standard error. In Column 2, we add the UPC effects, the R^2 increases considerably and the tax coefficient is still 0.99 when rounded to two digits. In this detailed model where we can explain 94 percent of the product/state/year variation in prices for a good over a six year period, we find that 99 cents of a $1.00 tax hike are passed on in the form of higher taxes and we cannot reject the null this coefficient is $1.00.

The results in the this section indicate that although excises taxes reduce the price of higher quality goods, the market share of lower quality goods increases after tax hikes, a result in contradiction with the basic model of Barzel. There are a number of possible explanations for why the flight to quality does not occur in this case. One possible explanation is that the markets for discounts and name-brands may be catering to different clientele with little between-brand switching in response to tax hikes. The results in Table 6 could be generated by a situation where smokers respond to tax hikes by reducing or quitting smoking and the demand elasticity is larger in absolute value for name-brand than discount cigarettes. This does not seem likely since it has generally been assumed that name brand cigarettes are household names because people are more attached to those brands and hence, these products should have a lower elasticity of demand. More importantly, Cummings et al. (1997) demonstrate that generic cigarette use is much heavier among lower income smokers and a number of authors demonstrate that the elasticity of demand for cigarettes is larger in absolute value for low income compared to higher income groups (Evans and Farrelly, 1998; Farrelly et al., 2001). More likely, this study corroborates the results in Hyland et al. (2005) who find that smokers are much more likely to self-report a switch to generics when faced with higher cigarette prices.

Excises taxes and the method of purchase

If brand loyalty is strong and between-brand shifting is small, smokers may respond to tax hikes by making different intra-brand purchase decision that signal a flight to quality. Cigarettes can be purchased from different sources (e.g., supermarkets, gas stations, convenience stores, superstores, online, etc.) and in a variety of sizes, with the two most frequent being by pack (20 cigarettes) or a carton (10 packs). The unit price of cigarettes varies considerably based on where it is sold and how it is sold. Per pack prices are substantially higher at convenience stores compared to other outlets and for packs versus cartons.

Given the data we have, we cannot examine the between-store movement in sales (e.g., supermarkets to convenience stores) but we can examine whether cigarettes are more likely to be sold as single packs or cartons.

Although the cigarette is the same whether sold as a pack or carton, the price difference reflects an aspect of quality, namely convenience. Cartons require storage and require a greater commitment to continue smoking. Packs on the other hand can easily fit in a shirt or pants pocket or a purse. Cartons also require a greater outlay of cash at the time of purchase.

Descriptive statistics about carton versus per pack sales are reported in the bottom half of Table . In this case, sales of single pack on average are about 12 percent more expensive than per pack costs for carton sales ($3.89 versus $3.47) and single pack sales in supermarkets represent about 55 percent of the market.

As with our analysis of generic and name-brand cigarettes, we construct a data set that varies by state, month, year and type of sale (carton versus pack) for a total of 4176 observations (29*12*6*2). Initially we present a graphical analysis of the impact of changing excise taxes on retail prices for carton and pack sales by looking at data for the same six states considered in Figures 1 and 2 (Arizona, Massachusetts, Ohio, Washington, Pennsylvania and Connecticut). In Figure 3, we report the average retail price for pack sales (the solid line) and carton sales (dotted line) before and after tax hikes. These figures indicate that the tax change generates immediate price hikes in most states, and the absolute price changes for pack and carton sales are virtually identical.

In Figure 4, we consider the market share for carton sales and unlike generics where we saw no flight to quality, in five of six states, we do see changes in carton market shares immediately after the tax hikes. The absolute change in carton sales market share is also very large. The tax hikes in Massachusetts, Ohio, Pennsylvania and Washington, ranging in size from 60 to 80 cents per pack, produced about a 5 percentage point drop in carton sales market shares while the 60¢ hike in Connecticut produced a 9 percentage point drop.

Initially, we examine how higher excise taxes impact retail prices for these sectors by estimating a model similar to that in equation 1. In this model the outcome of interest is the average per pack price for a carton or single pack in a state/month/year sale. These results are reported in Table 8. In each model, we include fixed state and month/year effects plus a dummy variable for sales by carton. Observations are weighted by packs sold and we allow for arbitrary correlation in errors within a state over time. In model 1, we regress total retail price on the real tax and estimate that prices rise by 97 cents for a $1.00 increase in per-pack taxes. We cannot reject the null that the coefficient on taxes is 1 in this context. In this model, we find that average price per pack is almost 50 cents lower for carton purchases.

In model (2), we allow the tax effect to vary based on carton and single pack sales. Here we find that carton sales increase by $1.03 per pack for a $1.00 per pack increase in the excise tax while the value is about $0.92 for single pack sales. As the p-values in the table indicate, we cannot reject the null that each coefficient is separately equal to 1 and we cannot reject the null that the coefficients are equal.

Using the parameter estimates in model (2), the results are clear that the relative price of per pack sales falls given a tax increase. If per pack prices for carton sales are $3.40 and tax rates increase by $0.50 per pack, then retail prices would rise 15 percent to $3.92/pack. In model (2) we see single pack prices are $0.61 higher than per pack carton prices and a $0.50 tax would raise retail prices by $0.47, increasing prices from $4.02 to $4.49. Prior to the tax hike, single pack prices were 18.2 percent greater than carton sale prices per pack. After the tax hike, this difference is now 14.5 percent. In contrast to the results for generics, we find a flight to quality from tax hikes in this context.

Next, we investigate how excise taxes impact carton market share by estimating a model similar to equation 3. For model (3) in Table 8, we regress the market share of packs sold by carton on fixed state and month/year effects plus the real excise tax per pack. The coefficient is a statistically significant -0.062 indicating that a $0.50 per-pack tax hike would reduce the carton market share by 3.10 percentage points.

The results in Table 6 indicate little between-brand substitution as a result of higher cigarette excise taxes while the results in Table 8 are suggestive of large within-brand changes in the way cigarettes are purchased. The results in Table 8 could however be due to some between-brand substitution if carton purchasers stop buying their current brand and buy a new brand by the pack. To examine whether this is within- or between-brand substitution in the form of purchase, we re-aggregate the data to the brand/purchase type/state/month/year cell. In this case we allow brands to vary by the type of cigarette, so, Winston, Winston Lights, Winston Ultra Lights, Winston Select and Winston Select Lights are all separate brands. In this case, the data now varies by brand (b) and by carton or single pack purchase (indexed by i), and the model can be written as

$$(5) \quad P_{bist} = \beta_0 + Tax_{st}\beta_1 + Multitax_{st} * Tax_{st}\beta_2 + CARTON_i\beta_3 + \eta_{1b} + \lambda_{1s} + \phi_{1t} + \varepsilon_{1bimt}$$

The only new terms in the equation are the inclusion of a brand fixed effect η_{1b} which captures permanent differences in prices attributed to brands and a dummy variable for whether the observation represents a sale by carton or not.

With sales in all states for all months, each brand would have a maximum of 4176 (29*12*6*2) observations. There are 358 brands so the maximum number of observations for equation (5) is 1.5 million. However, many of the smaller brands have no sales in many states so the number of observations is reduced to about 361,000.

Estimates of equation (5) are presented in the first column of Table 9. As before, observations are weighted by packs of sales, the model allows for arbitrary correlation in errors within a state, and the model includes fixed state, month/year and brand effects. The estimates for model (1) are similar to those in the first column of Table 8 in that we estimate an excise tax increase of $1.00 increases retail prices by $0.99, but we cannot reject the null the true coefficient is 1. In model (2) we allow the tax effect on prices to vary for pack and carton sales, and the results are again very similar to results when data is aggregated to the state level. Specifically, a $1.00 excise tax hike raises retail prices for pack and carton sales by $0.96 and $1.05, respectively, but we cannot reject the null that these estimates are the same.

In model (3), we estimate a model similar to that represented by equation (3) where we examine the impact of higher tax rates on carton sales. In this case, the dependent variable is the market share for carton sales among brand b in state s in period t and the dependent variable is defined as MS^c_{bst}. The estimating equation can be written as

(6) $MS^{c}_{bst} = \alpha_0 + Tax_{st}\alpha_1 + Multitax_{st} * Tax_{st}\alpha_2 + \eta_{1b} + \lambda_{2s} + \phi_{2t} + \varepsilon_{bst}$

where all terms are defined as above. When we look at the brand-specific market share for carton sales, the coefficient on real taxes is a statistically significant -0.064 and very similar to the -0.062 estimate in Table 8. The similarity in the estimates between the two tables for the market share equations suggests that all of the flight to quality in Table 8 is intra-brand shifts in the types of purchases.

Conclusion

The results of this study contribute to two different literatures. The first is a group of papers that consider the price effects of changing excise taxes. Using monthly scanner data on cigarette sales in supermarkets in 29 states, we estimate that retail prices increase dollar for dollar with excise tax changes. The impact of taxes on prices is the same for name-brand and discount cigarettes, and in a sample that includes 1.12 million observations at the state, month and UPC level, our estimate is that for a $1.00 tax hike retail prices increase by $0.99. These results suggest that smokers pay most of the tax burden of higher excise taxes.

We also examine whether excise tax hikes generate a flight to quality *a la* Barzel (1976). In contrast to some previous work, we find no tax-induced shift to name-brand cigarettes, whose price has fallen relative to cheaper brands. If anything, our results point to a small positive movement towards generic varieties although this shift is both quantitatively small and statistically insignificant. In contrast, we find a much larger shift in the way cigarettes are purchased with the market share of single pack sales increasing considerably after tax hikes.

Results from models with brand-specific data suggest this movement is almost exclusively within brand and not between brands, suggesting a fair degree of consumer loyalty.

Our work expands on papers in these two literatures by exploiting scanner data on individual products for a large cross section of stores in a large number of states. The time period is also ideal in that we have a number of large tax changes to exploit in the econometric model.

Our paper does have some limitation. Our paper illustrates the benefits of scanner data but our sample is from one retail line of business—supermarkets. We observe one potential shift pertaining to carton versus pack purchases, but there are many other possible shifts in purchase patterns such as across stores (e.g., gas station/convenience store purchases versus supermarkets).

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Figure 1
Nominal Retail Prices for Name and Discount Brand Cigarettes,
Before and After Major Excise Tax Hikes in Six States, IRI Data

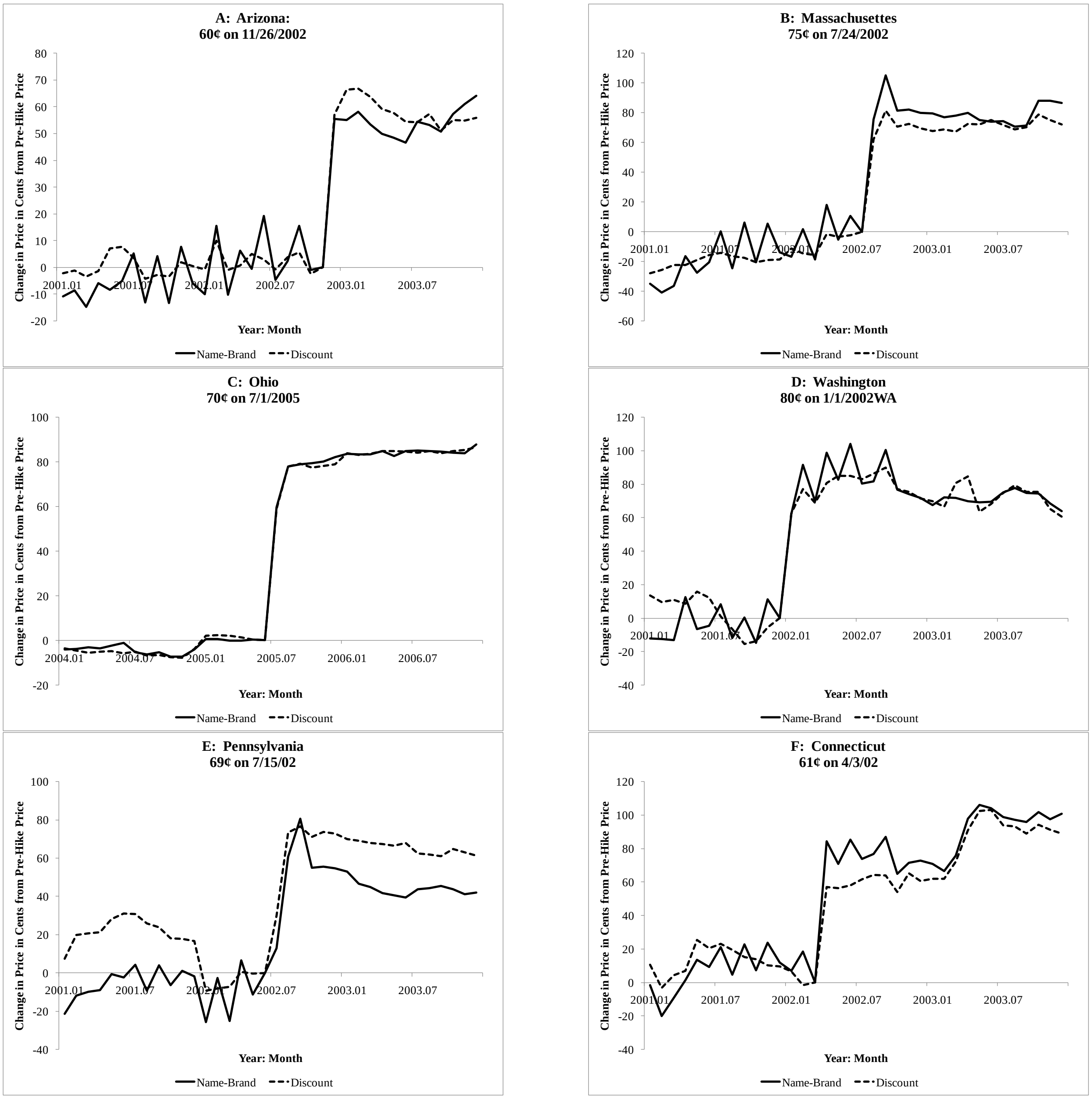

Figure 2
Discount Brand Market Share Before and After
Major Excise Tax Hikes in Six States, IRI Data

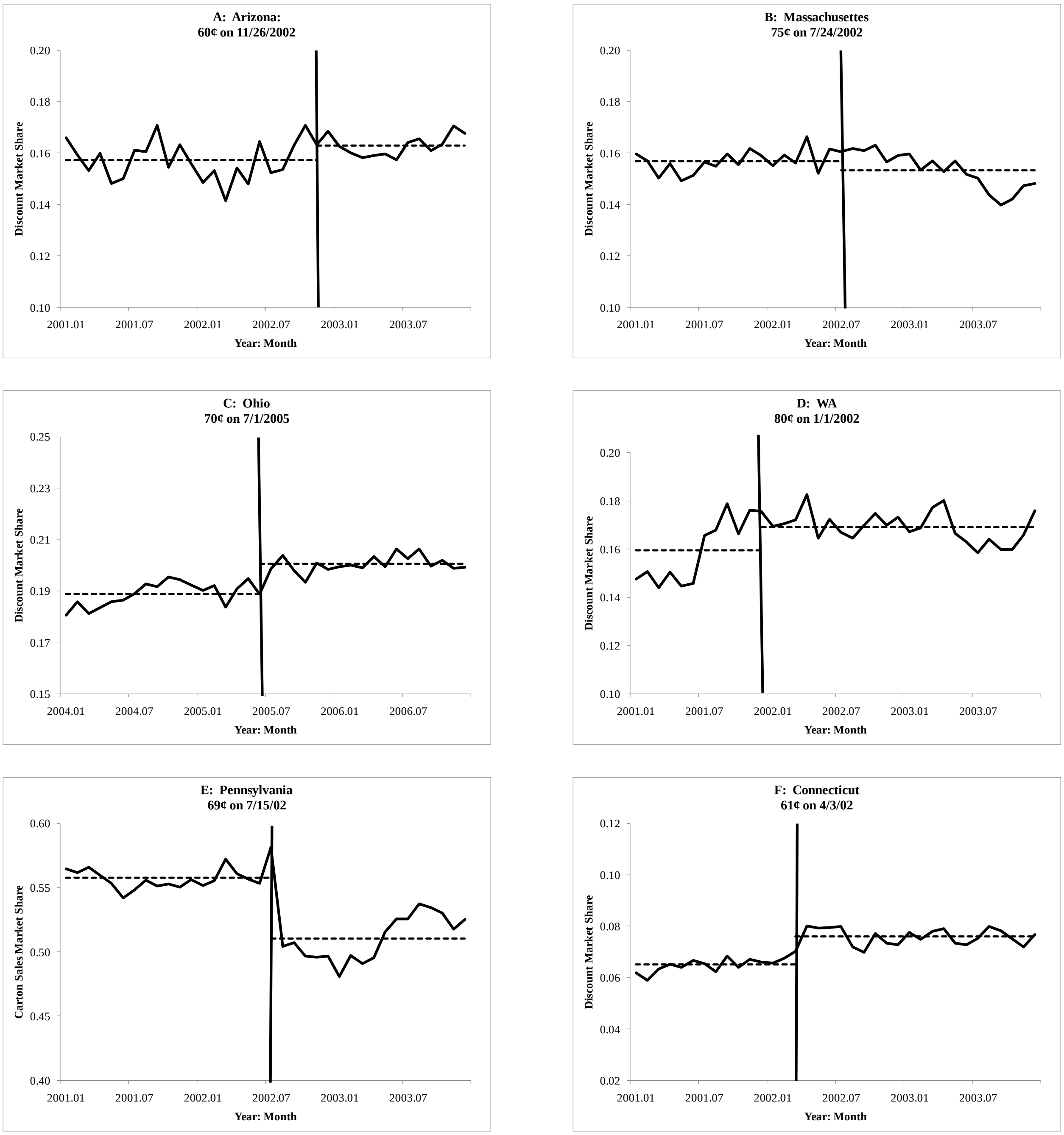

Figure 3
Nominal Retail Prices for Cigarettes for Pack and Carton Purchases,
Before and After Major Excise Tax Hikes in Six States, IRI Data

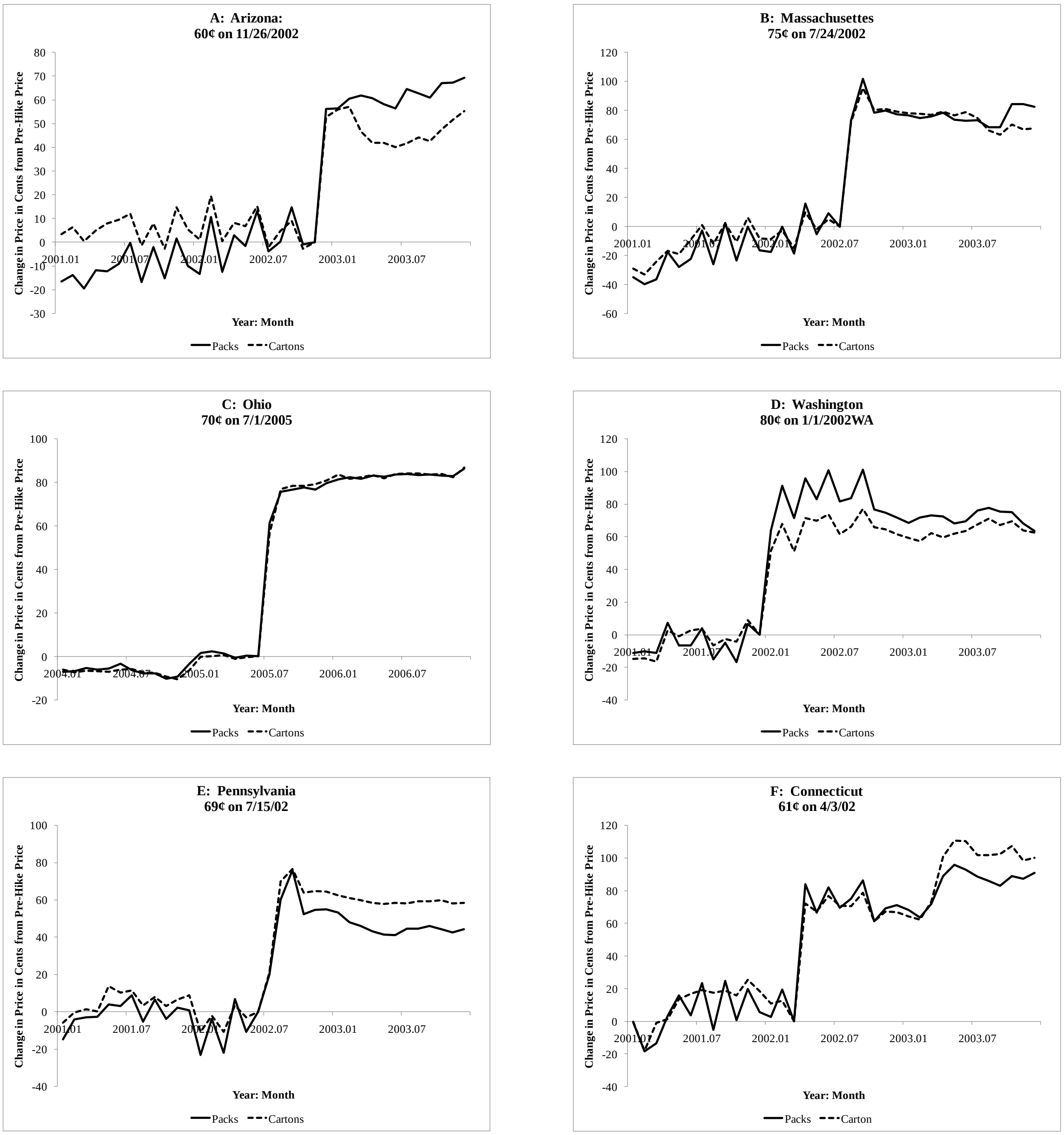

Figure 4
Carton Sales Market Share Before and After
Major Excise Tax Hikes in Six States, IRI Data

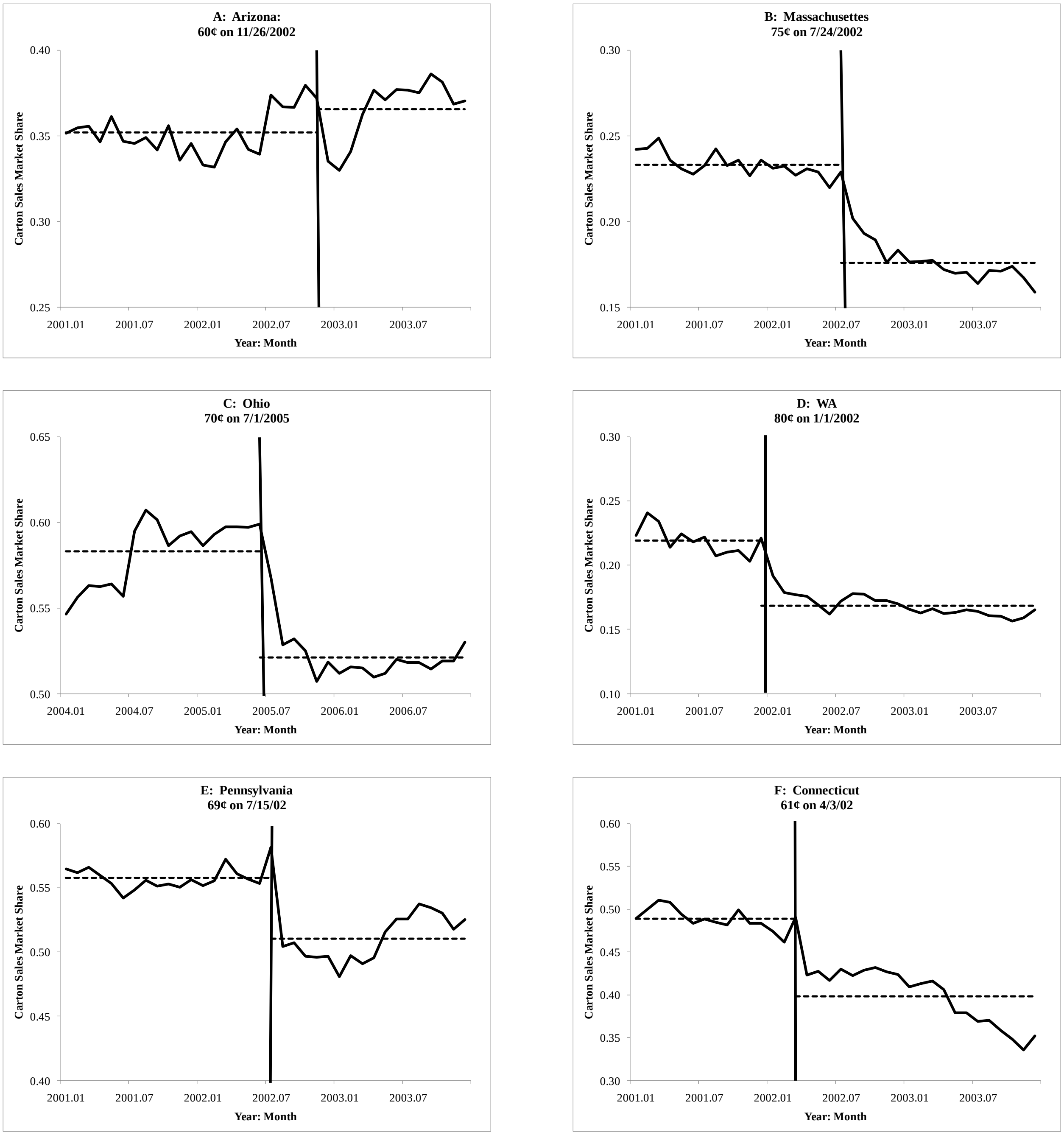

Table 1:
Markets and Stores in Sample, IRI Data

State	Markets	Stores	State	Markets	Stores
AL	Birmingham	16	NC	Charlotte	18
AZ	Phoenix	23		Raleigh/Durham	28
CA	Los Angeles	90	OH	Cleveland	11
	Sacramento	19		Toledo	9
	San Diego	23	OK	Oklahoma City	8
	San Francisco	40		Tulsa	6
CT	Hartford	19	OR	Portland	20
GA	Atlanta	28	PA	Harrisburg/Scranton	21
IA	Des Moines	7	RI	Providence	8
IL	Peoria/Springfield	6	SC		27
IN	Indianapolis	12	TN	Knoxville	12
LA	New Orleans	16	TX	Dallas	36
MA	Boston	33		Houston	23
	Pittsfield	14	UT	Salt Lake City	13
MI	Detroit	23	VA	Richmond/Norfolk	21
	Grand Rapids	12		Roanoke	19
MN	Minneapolis/St. Paul	12	WA	Seattle/Tacoma	24
MO	Kansas City	20		Spokane	4
	St. Louis	17	WI	Eau Claire	5
MS		6		Green Bay	6
NE	Omaha	9		Milwaukee	15
NY	Buffalo/Rochester	21			
	Syracuse	12			

The sample is restricted to store operating from January 1, 2001 through December 31, 2006. We exclude data from the New England, West Texas/New Mexico, Chicago, New York City and Philadelphia markets.

Table 2
State Tax Hikes within the Sample, 2001-2006

Year	State	Effective Date	Increase (Cents/pack)	Year	State	Effective Date	Increase (Cents/pack)
2001	RI	7/01/01	29¢	2003	CT	3/15/03	40¢
2002	AZ	11/26/02	60¢		GA	7/1/03	25¢
	CT	4/3/02	61¢		RI	7/1/03	39¢
	IL	7/1/02	40¢	2004	AL	5/18/04	29¢
	IN	7/1/02	40¢		MI	7/1/04	75¢
	LA	7/1/02	12¢		OR	1/1/04	-10¢
	MA	7/24/02	75¢		PA	1/7/04	35¢
	MI	8/1/02	50¢		RI	5/1/04	75¢
	NY	7/1/02	39¢		VA	9/1/04	17.5¢
	NE	10/1/02	30¢	2005	MN	8/1/05	75¢
	OH	7/1/02	31¢		NC	9/1/05	25¢
	OR	11/1/02	60¢		OH	7/1/05	70¢
	PA	7/15/02	69¢		OK	1/1/05	80¢
	RI	7/1/02	32¢		VA	7/1/05	10¢
	TN	7/15/02	8¢	2006	AZ	12/08/06	82¢
	UT	5/6/02	18¢		NC	7/1/06	5¢
	WA	1/1/02	60¢				

Source: Campaign for Tobacco Free Children, Research Center "Fact Sheets" on state tobacco taxes, http://www.tobaccofreekids.org/research/factsheets/pdf/0275.pdf
States without state tax changes within our sample include CA, IA, MO, MS, TN, TX, UT, and WI.

Table 3
Establishments selling tobacco products in the United States,
2002 Census of Retail Trade

NAICS Code Subsector	Industries	Industry description	Number of establishments selling tobacco products*	Sales of tobacco products ($1,000)	% of total Sales
445	44511	Supermarkets and other grocery (except convenience stores)	51,343	$7,683,683	15.1%
	44512	Convenience stores	27,871	$4,527,232	8.9%
	4453	Beer, wine, and liquor stores	13,177	$1,237,603	2.4%
446		Health and personal care stores (includes pharmacies and drug stores)	17,761	$1,525,046	3.0%
447	44711	Gasoline stations with convenience stores	86,152	$21,153,629	41.6%
	44719	Other gasoline stores	8,745	$1,019,700	2.0%
452		General merchandise stores (e.g., department stores, warehouse clubs)	6,991	$7,107,737	14.0%
453	453991	Tobacco stores	6,184	$5,674,466	11.2%
		Subtotal --	218,224	$49,929,096	98.2%
		Total for nation	221,173	$50,860,948	100.0%

The table includes all establishment types selling over $1 billion of tobacco products in 2002.

Table 4
Some Typical Name and Discount Cigarette Brands

Name-Brand		Discount	
Brand	Manufacturer	Brand	Manufacturer
Benson	Phillip Morris	Baileys	S&M Brands
Camel	RJ Reynolds	Basic	Phillip Morris
Capri	RJ Reynolds	Doral	RJ Reynolds
Carton	RJ Reynolds	GPC Approved	RJ Reynolds
Kent	Lorillard	Liggett Select	Liggett
Kool	RJ Reynolds	Malibu	Commonwealth
Marlboro	Phillip Morris	Misty	RJ Reynolds
Merit	Phillip Morris	Monarch	RJ Reynolds
Newport	Lorillard	Montclair	Commonwealth
Pall Mall	RJ Reynolds	Shield	Premier Marketing
Parliament	Phillip Morris	Tahoe	S&M Brands
Salem	RJ Reynolds	Ultra	Premier marketing
Vantage	RJ Reynolds	USA Gold	Commonwealth
Virginia Slims	Phillip Morris	USA	Vector
Winston	RJ Reynolds	Viceroy	RJ Reynolds

Table 5
Descriptive Statistics, IRI Data Set, 2001-2006

	Name-Brand	Discount
Name-Brand vs. Discount Sales		
Average real retail price/pack	$3.74	$3.13
Average real tax/pack	$1.04	$1.16
Average real before tax price/pack	$2.70	$1.97
Market share	81.7%	18.3%
Observations	2,088	2,088
	Sales by	Sales by
Pack vs. Carton Sales	Pack	Carton
Average real retail price/pack	$3.89	$3.47
Average real tax/pack	$1.16	$1.16
Average real before tax price/pack	$2.73	$2.31
Market share	54.7%	45.3%
Observations	2,088	2088

The unit of observation is the UPC/year/month cell. Prices are in real December 2006 prices.

Table 6
OLS Estimates of Real Retail Price and Discount Brand Market Share Equations, IRI Data 2001-2006

Parameter Estimates (Standard Errors) [P-value on Test H_o: β=1]

Covariates	Real Retail Price (1)	Real Retail Price (2)	Discount Brand Market share (3)
Real taxes ($/pack)	1.033		0.0205
	(0.086)		(0.0110)
	[0.702]		
Real taxes x Name Brand		1.038	
		(0.087)	
		[0.666]	
Real taxes x Discount Brand		1.013	
		(0.086)	
		[0.880]	
Discount brand	-0.736	-0.712	
	(0.028)	(0.056)	
R^2	0.978	0.978	0.942
Observations	4176	4176	2088
P-values:			
H_o: Real taxes x Discount = Real taxes x Name Brand		0.557	
H_o: Real taxes x Discount =1 and Real taxes x Name Brand=1		0.796	

There are 4176 observations in models 1 and 2 (29 states x 6 years x 12 months x 2 types of cigarettes) and 2088 observations in model 3 (29 states x 6 years x 12 months). All regressions control for state and month/year effects plus a control for months where observation contains data for two different state tax rates. Observations are weighted by the packs sold. Standard errors are calculated allowing for arbitrary correlation in errors within states.

Table 7
OLS Estimates of UPC-Level Real Retail Price Equation,
IRI Data 2001-2006

Covariate	(1)	(2)
Real tax ($/pack)	0.993	0.987
	(0.089)	(0.079)
	[0.938]	[0.872]
Month/year effects	Yes	Yes
State Effects	Yes	Yes
UPC effects	No	Yes
Observations	1,126,478	1,126,478
Distinct UPCs	2,843	2,843
R^2	0.687	0.939

Observations are weighted by the packs sold. The only over covariate in the model is the control for months with two different excise tax rates. Standard errors are calculated allowing for arbitrary correlation in errors within states.

Table 8
OLS Estimates of Real Retail Price and Carton Market Share Equations,
IRI Data, 2001-2006

Parameter Estimates (Standard Errors) [P-value on Test H_o: β=1]

Covariates	Real Retail Price (1)	Real Retail Price (2)	Market share of sales by carton (3)
Real taxes ($/pack)	0.965		-0.062
	(0.087)		(0.021)
	[0.687]		
Real taxes x Pack sale		0.924	
		(0.086)	
		[0.383]	
Real taxes x Carton sale		1.034	
		(0.107)	
		[0.753]	
Sold by carton	-0.497	-0.608	
	(0.050)	(0.081)	
R^2	0.968	0.969	0.962
Observations	4176	4176	2088
P-values:			
H_o: Real taxes x Pack Sales = Real taxes x Carton Sales		0.169	
H_o: Real taxes x Pack Sales=1 and Real taxes x Carton Sales=1		0.276	

There are 4176 observations in models 1 and 2 (29 states x 6 years x 12 months x 2 types of cigarettes) and 2088 observations in model 3 (29 states x 6 years x 12 months). All regressions control for state and month/year effects plus a control for months where observation contains data for two different tax rates. Observations are weighted by the packs sold. Standard errors are calculated allowing for arbitrary correlation in errors within states.

Table 9
OLS Estimates of Real Retail Price and Carton Market Share Equations,
IRI Data, 2001-2006

Parameter Estimates (Standard Errors) [P-value on Test H_o: β=1]

Covariates	Real Retail Price (1)	(2)	Market share of sales by carton (3)
Real taxes ($/pack)	0.991		-0.064
	(0.079)		(0.016)
	[0.904]		
Real taxes x Pack sale		0.960	
		(0.083)	
		[0.633]	
Real taxes x Carton sale		1.048	
		(0.088)	
		[0.588]	
Sold by carton	-0.464	-0.553	
	(0.043)	(0.071)	
R^2	0.937	0.938	0.808
No. of unique brands	358	358	358
Observations	360,677	360,377	160,291
P-values:			
H_o: Real taxes x Pack Sales = Real taxes x Carton Sales		0.212	
H_o: Real taxes x Pack Sales=1 and Real taxes x Carton Sales=1		0.284	

All regressions control for brand, state, and month/year effects, plus a control for months where observation contains data for two different tax rates. Observations are weighted by the packs sold. Standard errors are calculated allowing for arbitrary correlation in errors within states.